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Who has a Child as a Teenager?

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Further information about the BHPS and other longitudinal surveys can be obtained by telephoning +44 (0) 1206 873543.

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ABSTRACT AND NON-TECHNICAL SUMMARY

This paper uses data from the British Household Panel Survey (BHPS) and the British 1970 Cohort Study (BCS70) to investigate the family background and childhood factors that are associated with having a child as a teenager. The advantage of combining results from these two sets of data is that the BHPS analyses are restricted to a few background factors while the BCS70 analyses have far more. However, the results obtained from the BHPS data are reasonably replicated with the BCS70 data in that family social class and having lived with one parent during childhood are significantly associated with a higher likelihood of a teenage birth. From the BCS70 data we show that the effect of having lived with one parent is not significant once child-specific variables, such as self-esteem and teacher rated behaviour, are included in the models. Mother's age at the birth of the cohort member and mother's education have significant, consistent and robust associations with the likelihood of teenage birth.

The analyses reported in this paper are part of a larger programme of work for the Department of Health examining the medium and long-term consequences of early childbearing.

Who has a Child as a Teenager?

This question is examined with three sources of data. The first compares the impacts of family background factors on the chances that a woman has a child as a teenager from two sets of birth cohorts: women born during 1950-62 and women born between 1963 and 1976. The second examines the impact of the local economic environment, as well as family background, on the chances of a teen-birth over the 1990s, the women in question being born between 1976 and 1984. Finally, we use the substantial information on family background and childhood experience available for the 1970-birth cohort to study who becomes a mother as a teenager. Taken together, these three sets of analysis answer the question in the title for a range of different birth cohorts with data that allow different foci of analysis, and different measures of family background.

1. Trends in the impacts of family background

The data for this analysis come from the 1992 wave of the British Household Panel Survey (BHPS). Retrospective partnership and childbearing histories were collected from panel members, and these are linked to information on father's occupation when the woman was aged 14 (from the 1991 wave) and on whether she spent her entire childhood with both parents (from the 1996 wave). In order to examine trends in the impact of these two family background variables over time, we form two sets of birth cohorts: women born during 1950-62, who were primarily at risk of a teen-birth in the 1970s, and women born during 1963-76, who were mainly at risk in the 1980s. These two set of cohorts produce comparable sample sizes, although the later one has more 'censored information' (i.e. the 1992 interview intervened before they had a birth). Nevertheless, the statistical method employed uses all of the information efficiently.

The proportion of women becoming a mother as a teenager was similar in the two sets of cohorts: 13.1% for the 1950-62 cohorts, falling to 12.6% for the later cohorts.[1] The partnership context of these teenage first births did, however, change substantially, with the proportion married falling from 71% to 33%, and the proportion without a live-in partner at the time of the birth rising from 21% to 37%. A larger percentage of women came from a one-parent family in the later cohorts: 23%

compared with 15%. The other background factor examined, father's occupation when the woman was aged 14, is represented by the Hope-Goldthorpe score of the father's occupation. This score is strongly correlated with the earnings in that occupation.[2] Its mean value in the earlier and later sets of cohorts is 46 and 48, respectively, with standard deviations of about 15 in both sets.

Table 1: Percentage of Women becoming a Mother as a Teenager by Experience of a One-parent Family

Lived with both parents throughout childhood	**Percent**	**N**
No	18.2	380
Yes	9.5	1656

Table 2: Percentage of Women becoming a Mother as a Teenager by Father's Hope-Goldthorpe Score

Quartile of Father's HG Score	**Percent**	**N**
Bottom	22.0	373
3rd	11.8	414
2nd	12.1	504
Top	4.7	624
Missing information	11.5	615

Table 1 shows the percentage of women having a first birth as a teenager among all cohorts according to whether or not they had lived in a one-parent family sometime during their childhood, and Table 2 shows how this percentage varies with the quartile of the distribution of fathers' Hope-Goldthorpe (HG) scores.[3] The percentage is nearly twice as large for women from a one-parent family, and it is nearly twice as high for women whose fathers' occupation put them in the bottom quartile of HG scores as for women whose fathers were in the middle two quartiles. Furthermore, the proportion was much lower for women with fathers in the top quartile. Thus, these two family background factors are strongly associated with the chances of becoming a mother as a teenager, and so they appear to be good candidates for influences on the age distribution of the timing of first births.

There is assumed to be an underlying age pattern of the timing of motherhood, and different values of the two background factors speed it up or slow it down. In particular, the logarithm of the months between reaching age 14 and motherhood is assumed to have a Normal distribution with variance σ^2. This allows flexibility in the age pattern, depending on the estimated value of σ, and in particular it allows for the probability of becoming a mother at a given age among those still childless (i.e. the ‘hazard rate’ of a birth) to first increase and then decrease with age. In addition to the two background factors, each woman is characterised by an unobservable variable v that affects when she becomes a mother, and v is assumed to have a Gamma distribution with variance θ. One factor that this variable captures is fecundity—some women find it easier to conceive than others—and this is a likely to be a persistent influence for that woman. But it also captures volitional factors, including sexual activity. The model is presented formally in Appendix 1.

Two approaches to assessing the impact of the background factors on the chances of becoming a teen-mother are taken. First, we use all of the observed birth history information to model the distribution of age at first birth in the teenage years and beyond. From these estimates of the impacts of the background factors and the variance σ^2 we can calculate the probability that a woman has her first birth as a teenager for a woman with the mean value of the unobservable v. The second approach only models the age distribution of births over the teenage years; that is, it uses the exact birth timing information up to the age of 20, censoring all women who do not have a birth as a teenager at that age. The probability that a woman has her first birth as a teenager for a woman is calculated from the model estimates in a similar way.

The estimates of the parameters of the models are shown in Appendix 1. Model simulations of the probability that a woman has her first birth as a teenager for different family background factors are given in Table 3. Having experienced a one-parent family during childhood is associated with a substantially higher probability that a woman has her first birth as a teenager. Having a father in a lower earning occupation is also associated with a higher risk of a becoming a teenage mother. These associations are of a similar magnitude in both approaches (i.e. panels A and B), and they are also similar for the two sets of cohorts. In addition, there is evidence of persistent unobserved influences on the timing of motherhood from the model estimates using the entire observed distribution of age at first birth (leading to the simulations in panel A of Table 3). Because of the shorter observation interval, it is not possible to detect this ‘unobserved heterogeneity’ in first birth timing when women who are childless at age 20 are censored.

Table 3: Simulated Impact of Family Background on the Percentage of Women becoming a Mother as a Teenager

A. Based on entire observed first birth distribution

Family Background*	Birth Cohorts 1950-62	Birth Cohorts 1963-76
HGS=53, OPF=no	8.6%	7.1%
HGS=53, OPF=yes	14.8%	13.2%
HGS=33, OPF=yes	26.7%	23.9%
HGS=33, OPF=no	17.2%	14.3%

* HGS = Hope Goldthorpe Score
OPF = Lived with one parent sometime during childhood

B: Based on censoring of women who are childless at age 20.

Family Background*	Birth Cohorts 1950-62	Birth Cohorts 1963-76
HGS=53, OPF=no	7.0%	6.5%
HGS=53, OPF=yes	14.9%	12.6%
HGS=33, OPF=yes	28.4%	24.1%
HGS=33, OPF=no	15.7%	14.4%

* HGS = Hope Goldthorpe Score
OPF = Lived with one parent sometime during childhood

These results strongly suggest that women who become mothers as teenagers come from ‘poorer’ backgrounds, and this may affect them later in their life even if they had postponed motherhood.

2. Impacts of family background and the local economic environment in the 1990s

This section examines the impacts of family background and employment opportunities in local labour markets on the teenage first birth rate in Great Britain. It exploits variation in the unemployment rate in 300 “travel-to-work areas” over time and space to identify the latter impact. The data come from matching the first ten years of the British Household Panel Study (1991-2000) and the NOMIS (National On-line Manpower Information Service) travel-to-work area data. These data provide 2,367 woman-year observations on 849 childless women aged 16-19, who are at risk of a teenage first birth in the forthcoming year. Of the 84 teenage first births observed, 16 were born to women who were married, 29 to women in a cohabiting union and 39 to never-married women. Overall, the annual teenage first birth rate was 3.6 per cent. The rate was much higher for those who were in a cohabiting union in the previous year, 26.9% compared with 2.3% for those not in a live-in partnership, and 73% of women married in the previous year had their birth. Among women having a teenage first birth, one year after the birth 22% were married, 34% were in a cohabiting union and 44% were never-married.

As noted in the previous section, in the 1996 wave of the BHPS, all respondents were asked whether they lived with both natural parents up to the age of 16. Table 4 shows that young women who did not live with both parents throughout their childhood were more likely to have had a first birth as a teenager.

Table 4: Annual First Birth Rate as a Teenager by Experience of a One-parent Family

Lived with both parents throughout childhood	**Percent**	**N**
No	5.2	405
Yes	3.2	1445

Pearson chi2(1) = 3.6317 $p = 0.057$

These panel data allow us to observe each woman’s household income when living with her parents at age 16. Household income is converted into ‘equivalent income’ adjusted for household size by dividing it by the square root of household size. Table 5 indicates that the teenage first birth declines with the quartile of equivalent household income in the woman’s parental household at age 16.

Table 5: Annual First Birth Rate as a Teenager by Parental Household's Equivalent Income when aged 16-17

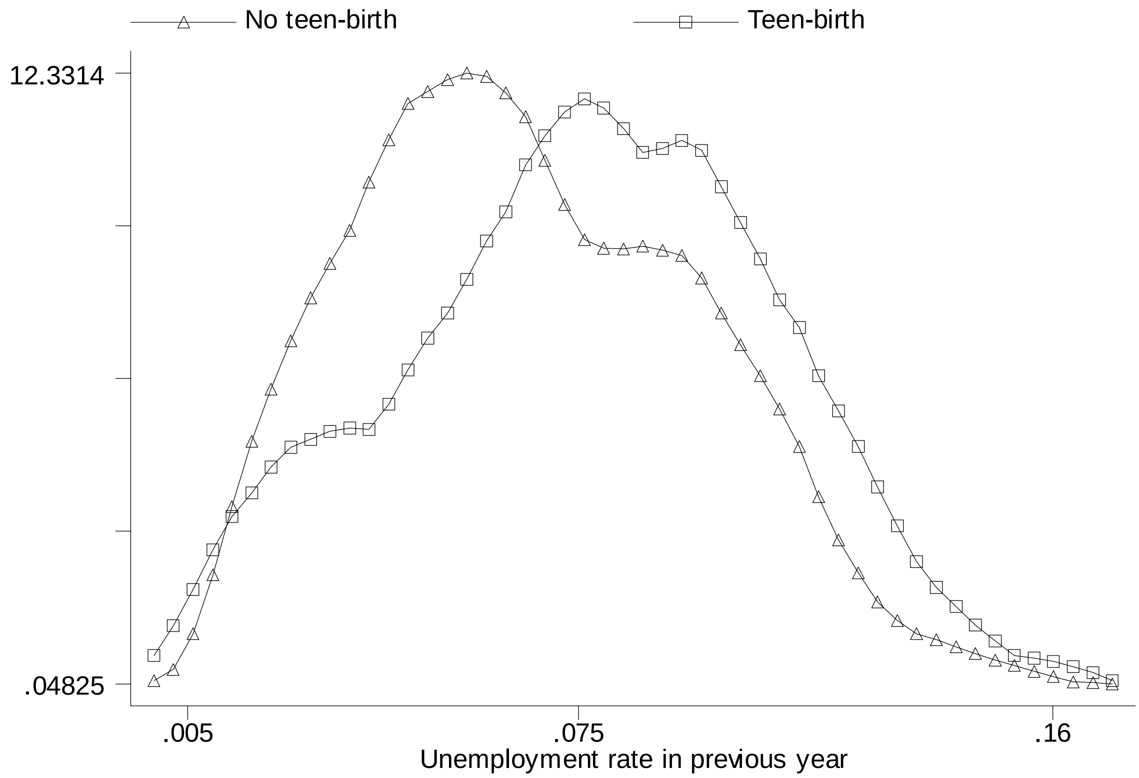
Quartile of HH Equiv. Income	Percent	N
Bottom	4.8	484
2nd	2.9	483
3rd	2.4	453
Top	1.2	482

Pearson chi2(3) = 11.15 *p* = 0.011

The chances of getting a job may also affect the decision to have a birth as a teenager. Poorer employment opportunities for young women reduce women's opportunity cost of childbearing, thereby tending to increase childbearing. This reasoning leads us to expect that women living in labour market areas with higher unemployment rates are more likely to have a child in the coming year. The mean local unemployment rate in the previous year is indeed higher for women who became a teen mother than for those who remain childless in the coming year, 7.6% compared with 6.9%, but an F-test indicates that this difference is not statistically significant at the 0.05 level (p=0.0566). When the sample is confined to women whose birth history we can observe from their 16th birthday onwards, then the difference in mean unemployment rates is statistically significant (p=0.017): 7.7% for women having a teen-birth compared with 6.6% for those who remained childless. Figure 1 compares the distribution of unemployment rates between those who had a teenage first birth in the following year and those who did not. The distribution among those who became mothers lies to the right, indicating that becoming a mother as a teenager was more common when local unemployment rates were higher.

We now estimate the impact of the unemployment rate on the age pattern of childbearing over the teenage years in a multivariate model that also contains the two background variables in Tables 4 and 5. By construction of the parental household income variable, the estimated models apply to those who we observe living with their parents at age 16. This reduces the sample size somewhat, but it also means that the sample is restricted to women whose birth history we can observe from their 16th birthday onwards.

Figure 1: Unemployment rates between those who had a teenage first birth in the following year and those who did not



The models we estimate are 'hazard rate' models; that is, they measure the influence of the local unemployment rate and the family background variables on the probability of becoming a mother at a given age, conditional on remaining childless up to that point. A more formal exposition of the model is given in Appendix 2. As we are interested in teenage births, we censor all women when they reach their 20th birthday. It turns out that once we control for parental equivalent household income (at age 16), whether or not a woman came from a one-parent family has no impact on the teenage first birth rate. This suggests that the significant association of this variable to teenage births in the previous section mainly reflects lower income in such families. Once parental family income is controlled for, there is no additional influence of an experience of lone parenthood as a child. Thus, only the results of models containing parental equivalent household income and the local unemployment rate in the previous year are discussed. As in the previous section, we use the model to simulate differences in the probability of having a teenage birth if the model applied to a cohort of women. The parameter estimates are given in Appendix 2.

Table 6: Simulated Percentage of Women becoming a Mother as a Teenager

Scenarios	Percent
Base Case*	9.8
One SD lower income	14.7
One SD higher income	6.5
Unemployment = 3.6%	7
Unemployment = 9.7%	13.5

*Mean log equivalent family income and mean unemployment rate (6.7%)

The base case in Table 6 indicates that the model predicts that 9.8% of women would become a mother as a teenager if the women's families had the mean level of equivalent family income when she was 16 and they lived in a local labour market with the mean unemployment rate.[4] If their families had equivalent income one standard deviation lower than the mean (putting them in the bottom quartile of the equivalent income distribution), then 14.7% would become teenage mothers. Living in a local labour market in which the unemployment rate was 9.7% (one standard deviation higher than the mean rate of 6.7%) would increase the percentage becoming teen-mothers to 13.5%. Thus, both family background in terms of income and local labour market conditions have major impacts on the likelihood of becoming a mother as a teenager.

It should, however, be noted that it might not be family income per se that affects the risk of a teen-birth, but rather a number of aspects of family background that are correlated with family resources. There are also reasons to question whether the unemployment rate in the local labour market is really independent of unobserved attributes of women that affect their childbearing patterns.

One reason that this may not be true is through sorting of people into labour markets. It is possible that unobservable attributes that affect childbearing decisions (such as "career motivation") may also affect migration decisions and therefore the type of labour market in which the women reside. For example, more career-oriented women may move to areas with better employment opportunities (a lower unemployment rate) and also be less likely to have a child. If so, the estimates of the impact of the unemployment rate on the pre-marital birth rate would be inconsistent and tend to overstate the impact. In our data, 12% of teenaged women moved between travel-to-work areas each year. There is evidence that movement between labour market areas was significantly related to the unemployment rate of her area in the previous year, with higher unemployment increasing the chancing of moving. But movement was not significantly influenced by whether she had a birth. Nor is the area unemployment rate significantly lower in the following year amongst those who moved between areas (after controlling for the local unemployment rate in the previous year and the average unemployment rate in the following year).

It is also possible that there are unmeasured aspects of local labour market areas that affect childbearing decisions and these are correlated with the local unemployment rate. The fact that these are wide areas defined by commuting patterns and not residential neighbourhoods makes it more credible to assume that this is not the case. Nevertheless, if we are to interpret the positive coefficient of the local unemployment rate on the teenage birth rate as affecting it through the opportunity cost of childbearing, some evidence is required that the local unemployment rate affects this cost. A woman's pay should be directly associated with her opportunity costs of childbearing. Estimates of fixed and random effects models of a teenaged, childless woman's own pay as a function of her age and the local unemployment rate indicate that the local unemployment rate has a negative and statistically significant effect on her pay (among workers).[5] Thus, we take the impact of the local unemployment rate as evidence that poorer employment opportunities increase the chances of a birth among teenaged women by reducing the opportunity cost of becoming a mother.

3. Family background and childhood variables in the 1970 cohort data

Data were used from birth, age 5 and age 10 to predict those women who became teenage mothers. Table 7 provides a summary of these measures. At all ages data were obtained from the parents of the cohort member (CM) about the CM themselves and their family. At age 10 teachers were also asked about the CM's behaviour at school.

Table 7: Summary of measures used in BCS70 analysis

Birth	**Age 5**	**Age 10**
CM	*CM*	*CM*
Low birth weight	Rutter child scale	Behavioural problems
	EPVT	Ever in care
	Intelligence rating	Early puberty
		Self esteem
		Teacher ratings
		Popularity with peers
		Number of friends
		Co-operation
		Negotiate behaviour
		Concentration
Parents	*Parents*	*Parents*
Mother smoked during pregnancy	Mother/child separated > 1 month	Ever lived with one parent
Mother's education	Mother's education	Mother's education
Age of mother		Residential mobility
Household social class	Household social class	Household social class
	Household social class changes	

Five of the teacher ratings of personality/behaviour were highly significant for prospective outcomes in adulthood but including all of them in the same regression model 'washed out' the individual effects through multicolinearity. Table 8 shows the first-order correlations between the five scales.

Scales (1) and (3) were reverse coded to correlate positively with the other scales and then all were standardised to a mean of zero and a standard deviation of one. A test of internal reliability was done (Cronbach's alpha) and was satisfactory for a five-item scale – 0.82. The five items were then scaled to create one standardised scale. This scale was used in its continuous form and also as a series of quintiles. This summary measure can then be used to control for personality/behaviour at age 10.

Table 8: First-order correlations (Pearson's *r*) between teacher ratings

		2	**3**	**4**	**5**
1	Concentration	-0.41	0.37	-0.39	-0.37
2	Popularity with peers	-	-0.84	0.64	0.37
3	Number of friends		-	-0.59	-0.33
4	Co-operation			-	0.49
5	Negotiate behaviour				-

We examine some changes in family social class between birth and age 10. Changes. Family social class variables at birth, age 5 and age 10 were dichotomised into high family class (non-manual occupations RGSC classes I, II and IIIN) and low family class (manual occupations RGSC classes IIIM, IV and V). Eight possible combinations as shown in Table 9.

Table 9: Combinations of family social class at birth, age 5, and age 10 and coding categories.

Family social class			
birth	**age 5**	**age 10**	**category**
low	low	low	cont. low
high	high	high	cont. high
high	low	low	drop
high	high	low	drop
low	high	low	drop*
low	high	high	rise
low	low	high	rise
high	low	high	rise*

* last change having priority

Number of teenage pregnancies and births

In the data collection at age 30, 5,790 women were interviewed and 3,670 reported having ever been/being pregnant. Of these, 844 women reported a pregnancy before age 20 (14.6% of the sample of women at age 30). Including multiple pregnancies before 20 a total of 1,002 pregnancies were reported. From these pregnancies, 582 women (10.0% of the sample of women at age 30[6]) had 664 live births.

Regression models

In this instance we use proportional hazard models to estimate the effects of the independent variables on the outcome – a live birth before age 20. Proportional hazard models estimate the rate of transition to a birth, that is the probability that a woman who has not had a birth will have a birth at a particular time. These models provide unbiased estimates even with a large proportion of censored cases, as in this case as only a minority of women had given birth before age 20. Cox's partial likelihood in which the baseline hazard is unspecified is used. In the results reported, age was measured in years from the woman's 10th birthday. Other models were estimated using age in months and discrete time analysis with unobserved heterogeneity. There was no significant persistent unobserved heterogeneity and there was no significant difference between age in years or in months. Results are presented in Table 10.

The age of the CM's mother at the time of birth is the only variable that is entered in all models. The effects are fairly robust in that CM's born to teenage or young adult (age 20-23) mothers are significantly more likely to have a teenage birth themselves – approximately two to two and a half times more likely. Household social class has similar effects regardless of when it was measured during childhood. At birth, age 5 and age 10 those CM's in non-manual households are half as likely to have a teenage birth. When changes in the household social class are examined (Model 5) the results show that those CMs in continuously high social class households have 70% lower odds of a teenage birth compared to those from households with continuously low social class. Those CMs in the other two categories – dropping and rising social class – are significantly lower than the continuously low but significantly higher than the continuously high.

The CM's mother's education, specifically no qualifications, doubles the likelihood of a teenage birth whenever the education level is measured. The CM's mother smoking during the pregnancy raises the likelihood of a teenage birth in Model 1 but this effect is non-significant in Models 4 and 5 that include later measures. The effect of the standardised EPVT score from age 5 remains significant in all subsequent models indicating that an increase of one standard deviation in the test score reduces the likelihood of a teenage birth by 20-30%. The teacher rating scale has a strong negative effect on the likelihood of a teenage birth in that each standard deviation above the mean reduces the likelihood by 30-40%. Self esteem at age 10 is also significant in Models 3, 4 and 5 with the likelihood of a teenage birth increasing by 7-8% for each additional point indicating lower self esteem.

Table 10: Proportional hazard models of teenage birth

Model	**1**	**2**	**3**	**4**	**5**
At birth					
Mother's age (1970)					
under 20	2.52***	2.11***	1.96***	2.41***	2.23***
20/23	151**	1.42*	1.30	1.56*	1.54*
24/27	*ref.*	*ref.*	*ref.*	*ref.*	*ref.*
28/31	0.88	1.02	0.76	1.07	1.07
over 31	1.11	1.24	1.10	1.40	1.43
High household social class	0.57***			-	-
Mother – no qualifications#	1.74***			-	-
Smoked during pregnancy	1.50***			1.09	1.07
LBW	1.08			1.11	1.10
At age 5					
High household social class		0.48***		-	-
Mother – no qualifications#		1.88***		-	-
EPVT		0.69***		0.77**	0.80**
Rutter scale		1.02		1.01	1.01
Separated parents >1 month		1.14		1.01	1.02
Rated < normal intelligence		1.13		0.82	0.94
At age 10					
High household social class			0.56***	0.49***	-
Mother – no qualifications#			2.01***	1.70***	1.55**
High residential mobility			1.27	1.27	1.31
Ever lived with one parent			1.27	1.17	1.10
Ever in care			0.71	1.02	0.57
Behavioural/emotional problems			1.13	1.05	1.05
Early puberty			0.97	0.95	0.97
Teacher rating scale			0.65***	0.71***	0.73**
Low self esteem			1.08***	1.09***	1.09***
Change in household class					
Cont. low					*ref.*
Cont. high					0.32***
Drop					0.65*
Rise					0.53***
χ^2 (df)	195(8)	225(10)	229(13)	200(19)	203(21)
Person/years	51296	39930	32795	26244	26244
Persons	5235	4069	3343	2672	2672
Births in estimation sample	515	379	319	237	237

* p<.05 ** p<.01 *** p<.001

4. Comparability of results

The two sets of data used in these analyses have different strengths. Specifically, the BCS70 data have numerous measures during childhood prior to teenage years but is weak on outcomes as these are only measured at age 30. Conversely, the BHPS only has two childhood measures (for the cohorts studied here) but numerous outcomes that have been measured repeatedly. Therefore, it is not possible to conduct many comparisons of the results but it is possible to look at the same background factors in the BCS70 as done with the BHPS – social class and single parent.

In the results in Table 10 the variable indicating any time spent with one parent is not statistically significant, probably because of the other background variables included in the model. However, in a model restricted to household social class and any time with one parent, the parameter estimate is significant – see Table 11. Table 12 presents the simulated impact of the two factors on the percentage of women who become a mother as a teenager[7] – equivalent to Table 3 Panel B. The Hope-Goldthorpe scores used in the simulations in Table 3 are the mean scores for those in the non-manual and manual occupations captured by the high/low household social class dichotomy. Both sets of results indicate that the family factors of social class and lived with one parent operate similarly in both the BHPS and BCS70 data.

Table 11: Proportional hazard results – restricted model

High household social class	0.39***
Ever lived with one parent	1.68***
χ^2 (df)	134(2)
Person/years	48463
Persons	4943
Births in estimation sample	484

* p<.05 ** p<.01 *** p<.001

Table 12: Predicted percentage of teenage births by household social class and ever lived with one parent

		Household social class	
		High	Low
Ever lived with one parent	Yes	8.8	21.1
	No	5.2	13.2

5. Conclusions

Family social class, whether measured using the Hope-Goldthorpe score or RGSC is a robust predictor of teenage births. Having lived with one parent is less robust with its effects becoming non-significant in the presence of other factors measured at age 10 – notably teacher's ratings of the child's behaviour/personality and an indicator of the child's self esteem. The age of the mother at the time birth and the mother's education have strong effects even after controlling for a host of child specific and family variables later in childhood. Those born to teenage mothers are twice as likely to become teenage mothers themselves. The child's development at age 5, measured by EPVT, also had a robust effect as did teacher's ratings and self esteem at age 10. However, none of these findings are surprising and generally reflect the findings from the NCDS and other cohort studies.

Appendix 1

In the absence of unobserved heterogeneity among women, the model estimated in part 1 takes the following form: $\ln(t_j) = \mathbf{x}_j\beta + z_j$, where t_j is the age at first birth for woman j, $\mathbf{x}_j$ is a vector of family background factors, β is a vector of parameters to be estimated and z_j has a Normal distribution with mean $\mathbf{x}_j\beta$ and variance σ^2. Age at birth, t_j, is measured as the number of months since a woman's 14^{th} birthday. The hazard rate of a first birth is given by $h(t_j)=\phi[(\ln(t_j)-\mathbf{x}_j\beta)/\sigma]/t\sigma\{1-\Phi[(\ln(t_j)-\mathbf{x}_j\beta)/\sigma]\}$, where $\phi[.]$ is the standard Normal density function and $\Phi[.]$ is the standard Normal distribution function. Unobservable heterogeneity is introduced as a multiplicative effect α_j on the hazard function: $h(t_j|\alpha_j)=\alpha_j h(t_j)$, where α_j has a mean of unity and a variance θ. In particular, α_j is distributed as Gamma$(1/\theta,\theta)$.

The two family background factors includes in $\mathbf{x}_j$ are the 'Hope-Goldthorpe Score' (HGS) of the occupation that the woman's father was in when she was aged 14, and whether or not she experienced life in a one-parent family during childhood. For women with missing information on the father's HGS score, we impute the mean value and an indicator variable for 'Missing HGS' is set equal to unity (zero otherwise).

Estimates of the parameters β, σ and θ for the two sets of birth cohorts are given in Table A1. The primary parameter estimates are similar in the two sets of cohorts. The large percentage of censored cases (66%) in the later cohorts makes it more difficult to estimate the unobserved heterogeneity distribution parameter, θ, leading to a relatively large standard error for the estimate of θ. Table A2 presents the parameter estimates when all childless women are censored at their twentieth birthday. Here it is impossible to identify the θ parameter.

The probability of becoming a teenage mother for a woman with the mean value of α_j is given by $\Phi[(\ln(72)-\mathbf{x}_j\beta)/\sigma]$. This is reported for particular values of $\mathbf{x}_j$ in Table 3.

Table A1

1950-62 Birth Cohorts

Parameter	**Estimate**	**Std. Error**	*P-value*
Father's HGS	0.0114	0.0014	0.000
Missing HGS	0.128	0.063	0.043
One parent family	-0.173	0.058	0.003
Constant	4.410	0.073	0.000
σ	0.537	0.023	0.000
θ	0.287	0.091	0.000

N=1,057; N of births=837; Chi-sq.(3)=72.53

1963-76 Birth Cohorts

Parameter	**Estimate**	**Std. Error**	*P-value*
Father's HGS	0.0110	0.0018	0.000
Missing HGS	-0.101	0.063	0.109
One parent family	-0.193	0.059	0.001
Constant	4.492	0.105	0.000
σ	0.542	0.038	0.000
θ	0.400	0.303	0.226

N=978; N of births=328; Chi-sq.(3)=53.47

Table A2

1950-62 Birth Cohorts

Parameter	**Estimate**	**Std. Error**	*P-value*
Father's HGS	0.0147	0.0028	0.000
Missing HGS	0.052	0.099	0.597
One parent family	-0.276	0.085	0.001
Constant	4.425	0.115	0.000
σ	0.629	0.048	0.000
θ	0	0.0054	0.986

N=1,057; N of (teen) births=125; Chi-sq.(3)=45.81

1963-76 Birth Cohorts

Parameter	**Estimate**	**Std. Error**	*P-value*
Father's HGS	0.0120	0.0027	0.000
Missing HGS	-0.154	0.070	0.027
One parent family	-0.195	0.068	0.004
Constant	4.451	0.117	0.000
σ	0.537	0.044	0.000
θ	0	0.067	0.990

N=978; N of (teen) births=101; Chi-sq.(3)=53.47

Appendix 2

Given that we are trying to model a births process, which is age-related, it is natural to model the first birth rate as a hazard rate, $h(a_i, \boldsymbol{X}_i, |v_i)$, where the population at risk is *childless* women, a_i is the age of the i-th woman, $\boldsymbol{X}_i$ is a vector of exogenous variables, including the local unemployment rate and family background and v_i is an unobservable individual effect on the non-marital birth rate that is distributed independently of a_i and $\boldsymbol{X}_i$. It is assumed that the hazard rate model takes the proportional hazard form; that is, $h(a_i, \boldsymbol{X}_i, |v_i)=\phi(a_i)\psi(\boldsymbol{X}_i)\nu_i$.[8] As we only measure the local unemployment rate annually, the discrete time representation of the model is used. This leads to the following specification of the model:

$$cloglog[h(a_{it}, \boldsymbol{X}_{it-1}|v_i)] = D_b(a_{it}) + \boldsymbol{X}_{it-1}\beta_b + ln(v_i) \qquad \text{(A1)}$$

where the *cloglog* transformation of h_{itj} is $ln[-ln(1-h_{itj})]$, which is known as the *complementary log-log* transformation; the t-subscript indicates when the variables are measured, a_{it} are discrete age intervals (years) and $D_b(a_{it})$ characterises the baseline hazard. The parameters in β_b measure the percentage change in the hazard rate from a unit change in an exogenous variable. The model is completed by specifying the baseline $D_b(a_{it})$ and a distribution for v_i. Identification of the parameters in β_b depends on assumptions about these, which are bound to be somewhat arbitrary. Three different restrictive assumptions are made, and the robustness of the estimate of β_b to different specifications is investigated. In one specification we assume that v_i has a Gamma distribution with variance σ^2. In a second, it is assumed that there are two types of women, one type having a zero risk of having a teen-birth and the other having a positive hazard but not additional unobserved heterogeneity within this type.[9] The third and most restrictive assumption is that v_i is a constant. All of these specifications assume a Weibull specification of duration dependence over the teenage years; that is, $D_b(a_{it})=ln(a_{it}-16)$. Because we are interested in teenage births, in the estimation all women are censored at the panel wave when they reach the age of 20.

We were not able to identify any significant unobserved heterogeneity. In the model in which v_i has a Gamma distribution, the estimated variance σ^2 is 0.62 with a standard error of 4.56. The estimates of the 'two-types', or 'split population' model implies that 45% of the women would never have a teenage birth, but the standard error of this estimate is 0.47. The failure to identify well the unobserved heterogeneity distribution probably reflects the limited observation period for teenage births. This was also evident when women's birth histories were censored at age 20 in the analysis in Table A2 of Appendix 1.

The estimates of the parameters β_b in the model in which v_i is a constant are shown in Table A3. The estimates of β_b in the two models with unobserved heterogeneity are similar to these. Empirical analysis suggests that the effects of unobserved heterogeneity on parameter estimates are reduced, making them more robust, if a flexible hazard baseline specification is used (e.g. see Dolton and van der Klaauw 1995). Estimates of a model in which $D_b(a_{it})$ is represented by an indicator variable for each year at risk as a teenager also produces similar estimates of β_b to those in Table A3, and there is little improvement in the log-likelihood relative to the model in Table A3 (log-likelihood=-214.37). Finally, Table A4 shows that the estimates of the impact of the unemployment rate and duration dependence are similar when equivalent family income at 16 is entered in terms of quartiles of its distribution rather than continuously (the reference category is the top quartile). The continuous specification in Table A3 performs better in terms of its log-likelihood.

Table A3

Parameter	**Estimate**	**Std. Error**	*P-value*
Unemployment rate, t-1	11.415	4.652	0.014
log(Equivalent family income at 16)	-0.741	0.207	0.000
log(age-16)	0.706	0.313	0.024
Constant	-0.1084	1.474	0.941

N=1,840; N of (teen) births=49; Chi-sq.(3)=22.71, Log-likelihood=-214.65.

Table A4

Parameter	**Estimate**	**Std. Error**	*P-value*
Unemployment rate, t-1	11.643	4.604	0.011
Equivalent family income, bottom quartile	1.414	0.544	0.009
Equivalent family income, 3rd quartile	1.039	0.578	0.072
Equivalent family income, 2nd quartile	0.924	0.585	0.114
log(age-16)	0.715	0.311	0.021
Constant	-6.1014	0.701	0.000

N=1,840; N of (teen) births=49; Chi-sq.(5)=20.47, Log-likelihood=-215.77.

Notes

[1] For comparison, birth registration statistics indicate that mean percentage of women in the 1950-62 birth cohorts who became mothers as teenagers was 16%, and it was 12% for the 1963-76 birth cohorts (see *Birth Statistics*, 1999, Table 10.3).
[2] In the BHPS, the correlation coefficient between gross monthly earnings and the Hope-Goldthorpe score was 0.7 during the 1990s.
[3] The substantially smaller sample in Table 4 reflects the fact that women who left the panel before the sixth wave would not have answered this question, nor would those who entered the panel after 1996.
[4] This is less than the 12% of women who become mothers born in 1975 indicated by official registration statistics; see *Birth Statistics*, 1999, Table 10.3.
[5] A Hausman test accepts the random effects model, and it indicates that each percentage point higher local unemployment reduces teenaged women's pay by 2.5% (t=3.15). The fixed effects estimator indicates a 4.8% reduction per one percentage point higher unemployment rate (t=2.68). A random effects model including equivalent family income at age 16 indicates an impact of -2.9% (t=3.58).
[6] Birth registration statistics indicate that 13% of the women in the 1970 birth cohort had a first live birth before their 20th birthday; see *Birth Statistics*, 1999, Table 10.3.
[7] Predicted percentages calculated from a logit model regressing teenage birth on the two family factors.
[8] Note that the form of the hazard rate assumed in Appendix 1 is not of the proportional hazard form.
[9] These two specifications are estimated by Stata programs *pgmhaz* and *spsurv* respectively. These were programmed by Stephen Jenkins.